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Comparing the Macroeconomic Effects of Conventional and Unconventional Monetary Policy in Japan

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Comparing the Macroeconomic Effects of Conventional and Unconventional Monetary Policy in Japan

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March 2015

Abstract

Using a sign-identified vector autoregressive (VAR) model, this paper compares the macroeconomic effects of a quantitative easing policy with those of conventional monetary policy in Japan. The results are as follows. First, I identify the significant macroeconomic effects of quantitative easing for the VAR specification in which expansionary monetary policy shock is defined as a shock that not only increases bank reserves but also compresses the long–short interest rate spread and output is allowed to respond to the monetary policy shock upon impact. Second, a qualitative comparison with conventional monetary policy suggests that quantitative easing has immediate, but less persistent, effects with greater uncertainty. Finally, assessing the quantitative differences from conventional monetary policy indicates that quantitative easing exerts stronger effects on output but weaker effects on the price level.

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1 Introduction

Many central banks in advanced economies adopted unconventional monetary policies in the wake of the 2007 global financial crisis. Recent studies have evaluated the macroeconomic effects of these policies, mainly using counterfactual simulation (e.g., Chung et al. (2012) for the US; Kapetanios et al. (2012) for the UK; Baumeister and Benati (2013) for both the US and the UK; and Lenza et al. (2010) for the Euro area). These all show that macroeconomic outcomes would have deteriorated further had the unconventional monetary policies not been implemented.\footnote{Using an estimated medium-scale dynamic stochastic general equilibrium (DSGE) model, Chen et al. (2012) found that the effects of the US Federal Reserve’s second large-scale asset purchase program on macroeconomic variables were likely to be modest, at least when compared with the above counterfactual analyses.} However, the counterfactual analysis conducted in their work is not a self-contained approach in that the impact of unconventional policy actions on a financial market variable, say long–short yield spreads, is taken as given and the value of the impact relies on estimates from other research or on direct observation of the data.\footnote{For example, in a US case analysis by Baumeister and Benati (2013), the assumption concerning the counterfactual path (the no-policy scenario) for the spread in 2009 is 60 basis points higher than the actual path (the policy scenario) according to the estimate in Gagnon et al. (2011).}

On the other hand, to evaluate the macroeconomic effects of unconventional monetary policy, some authors have employed a vector autoregressive (VAR) approach, which is a self-contained approach that has been most often used in the assessment of the effectiveness of monetary policy in more normal times (e.g., Christiano et al. (1999)). However, a relatively small sample period of unconventional monetary policies in the US, the UK, or the Euro area renders assessment with the VAR approach difficult. Gambacorta et al. (2014) attempt to address this problem by estimating a panel VAR model for eight advanced economies, but their sample period (2008–2011) remains to be short. Peersman (2011) applies the sign-identified VAR to the Euro area and identifies, as unconventional monetary policy shocks, innovations to the credit supply resulting from monetary policy actions that are orthogonal to a traditional interest rate policy. However, their sample period (1999–2009) includes the period of conventional monetary policy in the Euro area.

The Japanese experience of the early 2000s provides a good opportunity to estimate the macroeconomic effects of unconventional monetary policies, especially quantitative easing, based on a VAR approach with a relatively longer sample period. The Bank of Japan (BOJ) adopted quantitative

easing on three occasions: a first round from 2001M3 to 2006M3; a second round from 2010M10 to 2013M3 (entitled “Comprehensive Monetary Easing” by the BOJ); and a third round from 2013M4 onwards (referred to by the BOJ as “Quantitative and Qualitative Monetary Easing”). Indeed, there are many existing studies that have employed VAR-based approaches and investigated the macroeconomic effects of Japanese quantitative easing, including Kimura et al. (2003), Fujiwara (2006), Iwata and Wu (2006), Kamada and Sugo (2006), Inoue and Okimoto (2008), Franta (2011), Nakajima et al. (2011), Kimura and Nakajima (2013), Schenkelberg and Watzka (2013), Honda et al. (2013), Honda (2014), and Hayashi and Koeda (2014).

Unfortunately, these studies arrive at very different conclusions about the effectiveness of Japanese quantitative easing. Roughly speaking, significant effects on the macroeconomic variables are evident in Inoue and Okimoto (2008), Honda et al. (2013), Honda (2014), and Hayashi and Koeda (2014), while only transitory effects are evident in Franta (2011) and Schenkelberg and Watzka (2013), and very small, even insignificant effects are found in Kimura et al. (2003), Fujiwara (2006), Nakajima et al. (2011), and Kimura and Nakajima (2013).³

This paper revisits the assessment of Japanese quantitative easing policy, but the focus is on comparing the macroeconomic effects of quantitative easing with those of conventional monetary policy. To this end, I estimate a VAR model for the quantitative easing (QE) period from March 2001 to March 2014 and the normal period from January 1980 to June 1995. I then identify the monetary policy shocks as well as the money demand shocks along with two business cycle shocks by means of a combination of sign restrictions and magnitude and/or zero restrictions, and compare the impulse responses of the macroeconomic variables to the monetary policy shocks during the two periods.

The present analysis has three features not previously considered in the literature, each of which arises from the different natures of quantitative easing and traditional monetary policy. First, I estimate the VAR model for the QE period that includes the spread between long- and short-term interest rates, and define an expansionary monetary policy shock as a shock that not only

³Unfortunately, it is not possible to identify whether the effects are statistically significant in Iwata and Wu (2006) and Kamada and Sugo (2006) because of the lack of error bands for the impulse response functions. However, both studies show that at the level of the median (or average) response, the effects of a monetary policy shock on real economic activity become smaller when nominal interest rates are at their lower bound.
increases bank reserves, but also compresses the spread. This identification strategy is consistent with the view that direct purchases of longer-term securities under quantitative easing can decrease the term–premium component of longer-term interest rates.\textsuperscript{4} Fujiwara (2006), Kamada and Sugo (2006), Kimura and Nakajima (2013), and Schenkelberg and Watzka (2013) also include the long-term interest rate, although not the spread, in their VAR models, and show that the long-term interest rate falls in response to the expansionary monetary policy shock (at least in the short run and at the median estimate) during the low-interest-rate period. However, this paper \textit{a priori} assumes that an expansionary monetary policy shock decreases the spread as well as increasing bank reserves. Such an \textit{a priori} restriction should improve the accuracy of the identification of monetary policy shocks during the QE period.

Second, I allow for potential differences between quantitative easing and conventional monetary policy in terms of the times at which the effects of monetary policy arise. On the one hand, a zero lower bound on nominal interest rates may interrupt the transmission of quantitative easing. If this were significant, quantitative easing may affect the macroeconomy later than a more traditional monetary policy. On the other hand, quantitative easing may have more rapid effects on real economic activity if the aggregate supply curve is convex because of downward rigidity in nominal wages and prices.\textsuperscript{5} In addition, the price level may move up earlier in response to an expansionary monetary policy shock during the QE period than during the normal period, especially if the cost channel of monetary policy is smaller or disappears under the low-interest-rate environment.\textsuperscript{6} To consider these potential differences in the timing of the monetary policy effects, I adopt three identification variants (labeled Models I, II, and III) that differ in the strength of the restrictions on the responses of prices and output to monetary policy shocks. In brief, Model I allows both prices and output to immediately respond to monetary policy shocks, Model II constrains the initial response of prices to zero while allowing output to react immediately, and Model III imposes two

\textsuperscript{4}This view is based on the preferred-habitat model. See D’Amico et al. (2012) and D’Amico and King (2013) for discussion on the relationship between the effects of quantitative easing and the preferred-habitat model.

\textsuperscript{5}Relying on the hypothesis of a convex supply curve, Gambacorta et al. (2014) explain their empirical finding that unconventional monetary policy shocks have relatively larger output effects and smaller price effects than conventional monetary policy shocks. It is possible to use this hypothesis to explain the timing differences in the monetary policy effects between the two periods.

\textsuperscript{6}In the presence of the cost channel, an interest rate cut by the central bank will lead to a fall in the price level (i.e., a price puzzle) because it lowers the costs of firms associated with, say, working capital. See, e.g., Christiano et al. (1997) and Barth and Ramey (2001) for the cost channel of monetary policy.
zero restrictions on the initial responses of both prices and output.

Third, I conduct a quantitative comparison after equalizing the size of monetary policy shocks between the two periods. The literature usually reports impulse response functions to a one-standard-deviation monetary policy shock. However, in the context of the subject of this paper, there is a problem in that a one-standard-deviation monetary policy shock in the QE period is not equal to that in the normal period. This is because the main policy instrument of the BOJ changed from the call rate in the normal period to monetary aggregate measures, such as bank reserves or the monetary base, in the QE period. To address this, I estimate the size of a one-standard-deviation monetary policy shock in each period with a single measure that is common to both periods. In particular, I rely on information from the stock market. More specifically, I combine the VAR model with a stock return equation in which stock returns are regressed on all the structural shocks identified in the VAR. I can then regard the estimated response of stock returns to monetary policy shocks as the size of a one-standard-deviation monetary policy shock in a corresponding period that is comparable with both periods. Finally, I modify the impulse response functions for the QE period in such a way that they represent the responses to a monetary policy shock that are equal in size to those for the normal period.

The remainder of the paper is organized as follows. Section 2 describes the empirical method, Section 3 presents the results, and Section 4 concludes.

2 Empirical Methodology

2.1 Determining Sample Periods

To compare the macroeconomic effects of conventional and unconventional monetary policy, we first need to determine the two sample periods: namely, the normal period in which the BOJ manipulated the short-term interest rate as its policy instrument, and the QE period in which the BOJ adopted a quantitative easing policy. I define the normal period as that from January 1980 to June 1995. I select June 1995 as the end of the period because afterwards the call rate (the Japanese short-term interbank rate and the BOJ’s policy rate) fell below 1 percent (see the upper panel in Figure 1).
Meanwhile, the QE period is the period from March 2001 to March 2014, with the beginning point corresponding to the period in which the BOJ adopted the first round of quantitative easing.

Note that the QE period defined above contains intervals when the BOJ did not conduct quantitative easing but instead returned to the standard interest rate manipulation policy (from April 2006 to September 2010). Indeed, as shown in the lower panel of Figure 1, the amount of bank reserves declined rapidly after the termination of the first bout of quantitative easing, along with slight increases in the call rate. However, it remained above the required amount of reserves (i.e., there had been excess reserves) even during the period of nonexplicit quantitative easing. Moreover, bank reserves gradually increased in the aftermath of the global financial crisis.7 For these reasons, I regard this period as representing the implicit implementation of quantitative easing and assume that the changes in the amount of bank reserves in this period have the same effect on the economy as those in the period of explicit quantitative easing. In other words, we can interpret the estimated effect in this analysis as the average effect of explicit and nonexplicit quantitative easing over the QE period. By virtue of this assumption, I can estimate the effects of the quantitative easing policy with a sample size similar to that of the normal period.

2.2 The VAR Model

For each of the two periods, I estimate the following n-variable reduced-form VAR model:

\[ Y_t = c + B_1 Y_{t-1} + B_2 Y_{t-2} + \cdots + B_l Y_{t-l} + u_t, \]

where \( Y_t \) is an \( n \times 1 \) vector of endogenous variables, \( B_i \) is an \( n \times n \) matrix of autoregressive coefficients, \( c \) is an \( n \times 1 \) vector of constants, and \( u_t \) is an \( n \times 1 \) vector of reduced-form innovations that is normally and independently distributed with zero mean and the variance–covariance matrix \( \Sigma \). The lag length \( l \) is set to six months following Schenkelberg and Watzka (2013).

In the analysis of the QE period, I begin with a four-variable VAR model that consists of prices \( (p_t) \), output \( (x_t) \), the short-term interest rate \( (r_t^{\text{short}}) \), and bank reserves \( (m_t) \), which is common

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7The BOJ introduced a new funds-supplying operation in December 2009 to encourage a further decline in long-term interest rates.
to the normal period. However, while this specification is suitable for the normal period, it would not be appropriate in the QE period because the short-term interest rate near the zero lower bound includes less information on changes in the BOJ’s policy stance. Thus, as an alternative specification for the QE period, I also estimate a different version of the four-variable VAR that replaces the short-term interest rate with the spread \( r_t^{spread} \) between the long- and short-term interest rates. With this interest rate spread, I can more accurately identify the monetary policy shocks during the QE period, because one avenue for the expected transmission of quantitative easing is that direct purchases of longer-term securities can decrease the term–premium component of longer-term interest rates (see, e.g., D’Amico et al. (2012) and D’Amico and King (2013)).

For data on \( p_t \), \( x_t \), and \( m_t \), I use the consumer price index (all items less fresh food, seasonally adjusted), the industrial production index (seasonally adjusted), and the outstanding amount of bank reserves (adjusted for reserve requirement rate changes, seasonally adjusted), respectively. These are transformed into logarithms and multiplied by 100.\(^8\) As a measure of \( r_t^{short} \) I use the call rate (the Japanese short-term interbank rate and the BOJ’s policy rate), and as a measure of \( r_t^{spread} \) I use the difference between the 10-year yield on Japanese government bonds and the call rate. Appendix A provides a detailed description of the data.

### 2.3 Structural Shocks and Identifying Sign Restrictions

I consider the following four structural shocks: a price shock \( \epsilon_{pt} \), an output shock \( \epsilon_{xt} \), a money demand shock \( \epsilon_{mtd} \), and a monetary policy shock \( \epsilon_{mp} \). These are identified with a sign-restriction approach, which was developed by Uhlig (2005) and has often been employed in recent studies.\(^9\) In sum, this approach imposes some sign restrictions on the impulse responses, but leaves the responses of most interest to researchers unrestricted and allows them to be determined by the data.

The left panel of Table 1 presents the identifying sign restrictions proposed in this study. Note that I impose the same set of sign restrictions for the three VAR models (a VAR model for the normal period and two VAR models for the QE period, one with the call rate and the other with the

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\(^8\)I estimate the VARs in levels following Peersman (2011) and Gambacorta et al. (2014). For their analysis of Japanese quantitative easing, Schenkelberg and Watzka (2013), Honda et al. (2013), and Honda (2014) also estimate the VARs in levels, while Schenkelberg and Watzka (2013) further linearly detrend all variables.

\(^9\)See Peersman (2005) for the advantages of the sign-restriction approach over conventional strategies based on zero short- or long-run restrictions.
Also shown in parentheses are the periods over which I impose these sign restrictions, which I assume differ among the variables following Peersman (2005). For prices and output, I impose sign restrictions on the initial responses and subsequent five-month responses after a structural shock. For the short-term interest rate, the spread, and the bank reserves, I only impose the sign restrictions on the initial responses, as they are either a financial variable that can respond very quickly to a structural shock or a policy instrument variable that the BOJ can manipulate within a month.

For the VAR model with the call rate (irrespective of whether it is the QE period or the normal period), an expansionary monetary policy shock is defined as a shock that increases bank reserves and lowers the call rate at impact, following a standard money market model.\textsuperscript{10} For the VAR model with the spread, the sign restrictions are not changed, and an expansionary monetary policy shock is a shock that not only increases bank reserves but also compresses the spread. As discussed, this identification strategy is consistent with one of the most likely transmission mechanisms for quantitative easing.

Other than the monetary policy shocks, I also identify money demand shocks and two business cycle shocks. Once again, following the money market model, I distinguish money demand shocks from monetary policy shocks by assuming that a money demand shock moves bank reserves and the call rate (or the spread\textsuperscript{11}) in the same direction. Although identifying money demand shocks is redundant for the identification of monetary policy shocks, it would help to increase the probability of choosing true draws using a Bayesian estimation procedure, as described below.

As the business cycle shock, I identify price and output shocks in this paper. I assume that negative price and output shocks decrease prices and output, respectively. In addition, I assume that the central bank eases its monetary policy in response to such negative price and output shocks. I do not identify more primitive business cycle shocks, i.e., aggregate demand and aggregate supply shocks, in order to avoid considering a potential issue based on a prediction from Eggertsson’s\textsuperscript{12}

\textsuperscript{10}Uhlig (2005) also assumes this negative relationship between the policy rate (the federal funds rate) and the reserves (nonborrowed reserves) in addition to the sign restriction on the price response. In the literature on the Japanese case, Franta (2011) employs this negative relationship and further imposes the lower bound restriction on the call rate response, as well as the sign restriction on the output response.

\textsuperscript{11}For the VAR model with the spread, it is implicitly assumed that the spread (approximately equal to the long-term interest rate given that the short-term rate is stuck at a very low level) is the opportunity cost of holding money.
(2010) model, where the aggregate demand curve becomes upward sloping when the zero bound on
the nominal interest rate is binding. If this model is true, then the aggregate supply shock would
move prices and output in the same direction as an aggregate demand shock. However, because
it is unclear whether the aggregate demand curve has actually been upward sloping in the Japanese
economy during the QE period, I identify price and output shocks instead of aggregate demand and
supply shocks.

2.4 Further Restrictions on the Impulse Responses

The sign restrictions presented above are not sufficient to identify monetary policy shocks because
a price or output shock may generate a set of impulse response functions with the same pattern
of signs as a monetary policy shock. Consequently, I impose additional restrictions on the impact
responses of prices and output to monetary policy shocks in the form of a magnitude and/or zero
restriction.

In particular, three variants of identification are considered. The first identification (Model I)
imposes two magnitude restrictions whereby the size of the impact response of prices (output) to a
monetary policy shock is smaller than that of prices (output) to a price shock (an output shock).13
This identification strategy allows prices and output to respond to the monetary policy shock at
impact.

In the second identification (Model II), by applying the method used in Baumeister and Benati
(2013), I impose a zero restriction on the impact response of prices while a magnitude restriction
is placed on the impact response of output. This identification embodies the conventional wisdom
that the motion of prices is stickier than real economic activity.

Drawing on the approach in Gambacorta et al. (2014), the third identification (Model III) places
zero restrictions on the impact responses of both prices and output. Although the usual identification
scheme using Cholesky decomposition with a policy variable ordered after macroeconomic variables
also generates zero impact responses of both prices and output, the benefit of this approach is

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12 Taking into account Eggertsson’s (2010) model, Schenkelberg and Watzka (2013) identify aggregate supply and
aggregate demand shocks by imposing additional restrictions on the ratio of the output response to the price response.
13 Peersman (2005) imposes a magnitude restriction along with sign restrictions to distinguish between supply shocks
and oil price shocks.
that I can preserve the sign restrictions provided in Table 1, except for one minor restriction. See Appendix B for details of the identifications in Models I to III.

The three models have different degrees of strength in their restrictions, ranging from the weakest constraint in Model I to the strongest constraint in Model III. Therefore, they can capture potential differences in the time at which the effect of monetary policy arises between quantitative easing and conventional monetary policy, as discussed in the Introduction.

2.5 Stock-Returns Equation

One issue that arises when comparing monetary policy effects is that a one-standard-deviation monetary policy shock indicates different sizes between the normal and the QE period. This is because the main policy instrument has changed from the call rate in the former period to monetary aggregates such as bank reserves or the monetary base in the latter period. To address this issue, I estimate the size of a one-standard-deviation monetary policy shock in each period with a single measure common to both periods. In particular, I utilize information from the stock market and estimate the following stock return equation:

$$\Delta s_t = \beta_0 + \beta_1 \epsilon^p_t + \beta_2 \epsilon^x_t + \beta_3 \epsilon^{md}_t + \beta_4 \epsilon^{mp}_t + \eta_t,$$

(2)

where $\Delta s_t$ denotes stock returns, constructed by taking the log-difference (multiplied by 100) of the Japanese stock price index known as the Nikkei Stock Average, where the explanatory variables are four structural shocks identified from the VAR model, and $\eta_t$ is an iid disturbance term. Here, we can regard the contemporaneous response of stock returns to monetary policy shocks, $\beta_4$, as the size of a one-standard-deviation monetary policy shock in each period. I then adjust the impulse response functions estimated for the QE period such that they become impulse responses to an expansionary monetary policy shock that is equal in size to that of a one-standard-deviation shock in the normal period.

There are four points to be discussed with regard to Eq. (2). First, it is assumed that stock returns are determined outside of the VAR system (1). This is because the purpose of incorporating the stock return equation (2) is not to examine the transmission mechanism of monetary policy,
but merely to estimate the size of a one-standard-deviation monetary policy shock in each period. Second, it contains no lags of the structural shocks as explanatory variables because I consider stock returns as a fast-changing variable that responds immediately to new information. Third, the use of a financial variable to measure the relative size of monetary policy shocks requires the condition that the volatility of that financial variable should be the same for the two periods. This condition is satisfied in the case of Japanese stock returns with a standard deviation of 5.79 in the normal period and 5.86 in the QE period, suggesting no statistical difference. Finally, as presented in the right panel of Table 1, an additional sign restriction is imposed on Eq. (2) such that stock returns respond positively to an expansionary monetary policy within a month. This restriction is necessary for the estimated size of a one-standard-deviation monetary policy shock to be positive. Moreover, as a by-product of the additional sign restriction, the identification of monetary policy shocks can be improved because it means that an expansionary monetary policy shock that increases bank reserves and lowers the policy rate or the spread must also stimulate stock returns.

2.6 Estimation Procedure

Estimation of the VAR model (1) combined with the stock return equation (2) uses a Bayesian approach. I first draw for the VAR parameters from a Normal-Inverse Wishart posterior distribution, as well as drawing for the orthogonal matrix. I then compute the impulse response functions, and check whether they satisfy all the restrictions presented above. Finally, I estimate the stock return equations using OLS. A more detailed description of the estimation procedure is as follows:

Step 1. Take a random draw for the variance–covariance matrix Σ from the inverse Wishart posterior distribution, and given that draw, take a random draw for the VAR coefficients $B(= [c \ B_1 \ B_2 \cdots B_l]')$ from the normal posterior distribution. Then, I derive a lower triangular matrix $A$ by applying the Cholesky decomposition to Σ.

Step 2. Using the Cholesky factor $A$ as well as one or two orthogonal matrices, construct a structural impact matrix $\tilde{A}_I$, $\tilde{A}_{II}$ or $\tilde{A}_{III}$, depending on which model is estimated (see Appendix B).

14See Uhlig (2005) and Inoue and Kilian (2013) for more details about the Normal-Inverse Wishart (or the Normal-Wishart) posterior distribution of the VAR parameters. With regard to the hyperparameters of the Normal-Inverse Wishart prior distribution, I set the values following Uhlig (2005), i.e., using his notation, $N_0 = 0$, $v_0 = 0$, and $S_0$ and $B_0$ arbitrary (I actually set $S_0$ to a $4 \times 4$ identity matrix and $B_0 = 0$), which reduces to a noninformative prior.
Step 3. Using the VAR coefficients $B$ and the impact matrix $\bar{A}_I$, $\bar{A}_{II}$, or $\bar{A}_{III}$, compute the set of impulse response functions. If the impulse response functions satisfy all the restrictions, keep the draw. Otherwise, discard the draw.

Step 4. Using the VAR residuals and the impact matrix for the draw saved in the above step, calculate the time series of each of the four structural shocks.

Step 5. Using these structural shocks as explanatory variables, estimate the stock return equation (2) by OLS. If the coefficient on monetary policy shocks, $\beta_4$, is positive, keep the draw. Otherwise, discard the draw.

Step 6. Repeat the above steps until 1,000 sets of impulse response functions are obtained.

3 Results

3.1 Impulse Response Analysis for the QE Period

I first present the result obtained from the VAR model that consists of the call rate (the BOJ’s policy rate in normal times) as well as prices, output, and bank reserves, which is the same as the VAR model estimated for the normal period. Figure 2 plots the impulse response functions of these four variables to an expansionary monetary policy shock, with the columns corresponding to Models I to III. A solid line in each chart indicates the median responses to the shock. Although the VAR literature on the effect of monetary policy usually reports only the 16th and 84th percentile error bands, I also draw the 5th and 95th percentile error bands (dotted lines) in addition to the 16th and 84th percentile error bands (dashed lines). To facilitate comparison between Models I, II, and III, the size of the monetary policy shocks in Models II and III is adjusted such that the magnitude of the impact median response of bank reserves is equal to that in Model I.\textsuperscript{15} Note that I do not use the size adjustment of the monetary policy shock that relies on the stock return reactions as described in Section 2.5. In the next subsection I use this adjustment to compare the impulse responses between the QE and the normal periods.

\textsuperscript{15}In Model I, I report the impulse responses to the usual one-standard-deviation monetary policy shock.
As shown in Figure 2, an expansionary monetary policy shock in the QE period, which delivers a 4.5 percent increase in bank reserves and approximately a 1 basis point decline in the call rate at impact, has no significant effect on the price level and real economic activity. The exceptions are the price response at the fifth month in Model II and the output responses at horizons around the thirtieth month in Model III, which are only marginally significant based on the 16th and 84th percentile error bands.

However, as discussed previously, the short-term interest rate under a low-interest-rate environment tends to contain less information on changes in monetary policy because of its tiny fluctuations. Therefore, I use the spread between the long- and short-term interest rates instead of the call rate to improve the accuracy of the identification of the monetary policy shocks for the QE period. Figure 3 displays the impulse response functions obtained from the VAR model that replaces the call rate with the spread. Again, I equalize the size of the monetary policy shocks in Models II and III to that in Model I, which corresponds to a 4.9 percent increase in bank reserves. In Models I and II, this expansionary monetary shock lowers the long–short spread by about 2 basis points at impact, and stimulates both prices and output with significance based on the 90 percent error bands at horizons around a peak effect. Conversely, in Model III the responses of prices and output are reduced and significant only with the 68 percent error bands. Therefore, not only the use of the spread in the VAR but also weaker identifying restrictions on the impact responses of macroeconomic variables (Model I or II) are required to obtain highly significant estimates for the effect of quantitative easing in Japan.

A closer look reveals immediate and short-lived responses of output and prices regardless of the model. In response to an expansionary monetary policy shock, output immediately increases and reaches a peak in the fourth month in Models I and III and the second month in Model II, while prices display somewhat delayed increases with a peak in the ninth month in Model I and the eleventh month in Models II and III. The immediate response of output is also seen in Inoue and Okimoto (2008), but slower responses are found in Schenkelberg and Watzka (2013), Honda et al. (2013), and Honda (2014). The short-lived response of output is consistent with Franta (2011).

The short effect on output is in the ninth month in Honda et al. (2013), in the seventh month in Honda (2014), and around two years after the shock in Schenkelberg and Watzka (2013).
and Schenkelberg and Watzka (2013), but more persistent responses are detected in Inoue and Okimoto (2008), Honda et al. (2013), Honda (2014), and Hayashi and Koeda (2014). I can explain this short-lived effect of quantitative easing by the brief compression in the spread. As shown in the third line of Figure 3, the spread declines at impact, but returns to its original level within six months at the median estimate, and then turns to be positive. This implies that at the lower bound of the short-term interest rates, longer-term interest rates also decline for only a brief period after the expansionary policy shock. I explore this further in the next subsection, in which I calculate the implied impulse response function of long-term interest rates.

### 3.2 Comparison with the Normal Period

Next, I provide the result for the normal period and then compare it with that for the QE period. Figure 4 displays the impulse responses to an expansionary monetary policy shock, estimated using the normal period sample. As before, I adjust the size of the monetary policy shock in Models II and III such that the impact median response of the call rate is equal to that in Model I. This policy shock lowers the call rate by 26 basis points and increases bank reserves by about 1 percent for Models I and II and 1.6 percent for Model III at impact. In response to this shock, the price level initially declines (i.e., indicating a price puzzle), but gradually increases and finally exhibits significant and permanent responses. This response pattern of the price level in normal times is also evident in Inoue and Okimoto (2008) for the Japanese case and in studies on monetary policy effects in other countries (e.g., Christiano et al. (1999) for the US case). Traditional monetary policy shocks have a hump-shaped effect on real economic activity, which is in line with many previous studies. In particular, such effects are significant even with the 90 percent error bands in Models II and III. Note that this is in contrast to the result for the QE period, for which the most significant effect is found in Model I. This suggests that, when comparing the effects of conventional and unconventional monetary policy, it would be appropriate to consider several types of identifying restrictions on the impact responses of macroeconomic variables rather than a single common set.

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17 However, in Schenkelberg and Watzka (2013), the transitory effect on output is seen at later horizons (between 20 and 30 months after a quantitative easing shock).

18 However, some previous studies estimating the effects of conventional monetary policy in Japan with the VAR model in differences tend to find a permanent effect on real economic activity. See, e.g., Miyao (2000, 2002) and Inoue and Okimoto (2008).
I compare the monetary policy effects between the QE and the normal periods in more depth. As emphasized earlier, to make a quantitative comparison possible, I equalize the size of the monetary policy shocks between the two periods by relying on information from the stock market. Table 2 reports the median estimate of $\beta_4$ in Eq. (2) for the two periods and for Models I to III. The ratio of $\beta_4$ in the two periods, presented in the right column, is used as a scale adjustment factor to modify the impulse response functions for the QE period. A comparison of the monetary policy effects between the QE and the normal periods is shown in Figure 5, in which solid and dashed lines indicate the median and the 16th and 84th percentile error bands for the QE period (corresponding to Figure 3), while the shaded areas represent the regions within the 16th and 84th percentile error bands for the normal period (corresponding to Figure 4).\footnote{The impulse response functions for Models II and III in the normal period shown in Figure 5 are slightly different from those shown in Figure 4 because the former reverts to the original response to a one-standard-deviation monetary policy shock.}

There are several remarkable differences in the macroeconomic effects of monetary policy between the two periods. First, monetary policy shocks during the QE period have stronger effects on output and weaker effects on the price level compared with the effects of conventional monetary policy. More precisely, the peak effect on output measured by the median estimate is 1.1 percent in the QE period versus 0.6 percent in the normal period for Model I, 1.5 versus 0.7 percent for Model II, and 0.7 versus 0.5 percent for Model III. Meanwhile, the peak median effects on prices in the QE period (11, 13, and 8 basis points for Models I to III, respectively) are smaller than the long-run (at the 60th horizon) median effects in the normal period (14, 16, and 12 basis points for Models I to III, respectively). This quantitative difference is also evident in Gambacorta et al. (2014), in which they attribute their result to the convex shape of the aggregate supply curve from downward rigidity in nominal wages and prices. More concretely, if the aggregate supply curve is convex, an upward shift in the aggregate demand curve induced by monetary policy easing would lead to a larger increase in output and a smaller increase in the price level when an economy is at very low levels of both output and prices.

Second, uncertainty regarding the effect of quantitative easing is greater than that regarding the effect of traditional monetary policy. In particular, the evidence is seen in the output responses
rather than in the price responses. This reflects conflicting results among previous studies, argued in the Introduction, about the real economic effect of quantitative easing in Japan.

The third and fourth differences are that quantitative easing has more immediate but less persistent effects than traditional monetary policy, as pointed out in the previous subsection where there is only the result for the QE period. The output response to the monetary policy shock in the QE period reaches a peak at between two and four months, and the response becomes insignificant within a year based on the 68 percent error bands. In contrast, in normal times the output response is persistent with a peak at between one year and one and a half years. Similarly, a rise in the price level for the QE period is earlier and less persistent than that for the normal period; the former reaches a peak at between nine and eleven months and becomes insignificant within two years, whereas the latter shows a price puzzle in the short run but a significant permanent response after about two years.

Such immediate and temporary effects of quantitative easing may be specific to the Japanese case. Peersman (2011), who analyzes the effects of unconventional policy actions in the Euro area, shows that the effects of balance sheet policies on output and inflation are more sluggish than the effects of traditional monetary policy, which is contrary to the result in this paper. Using data from eight advanced economies, Gambacorta et al. (2014) find that output rises to a peak after about six months and then gradually returns to baseline after eighteen months, arguing that this response pattern is qualitatively very similar to the existing evidence on the transmission of conventional monetary policy shocks.20

20 On the other hand, Gambacorta et al. (2014) find that the price response to unconventional policy shocks is temporary, which is consistent with the result in this paper.

21 Because of a lack of data on the 10-year yield of Japanese government bonds in the normal period, I instead use the 9-year yield estimated and provided by the Ministry of Finance Japan.

Why are the macroeconomic effects of quantitative easing so short-lived, at least in Japan? One potential explanation is that the BOJ failed to compress the long-term interest rate over an extended period. Figure 6 displays the impulse responses to an expansionary monetary policy shock of one standard deviation, obtained from a five-variable VAR model that includes both the spread and the call rate, as well as prices, output, and bank reserves for the QE period (upper line) and the normal period (lower line), with zero restrictions on the impact responses of prices and output (i.e.,
Model III). In addition, the right column in this figure shows impulse responses of the implied long-term interest rate calculated by the impulse response of the spread plus that of the call rate. Although the error bands become slightly wider, especially for the normal period, the main results obtained above, i.e., more immediate but less persistent effects of quantitative easing, are preserved for this five-variable VAR. The important thing is that the falls in the implied long-term interest rates in response to an expansionary policy shock are transitory in the QE period, but those in the normal period are more long-lasting at the median estimate (although the latter negative responses are only significant until two or three months after the shock). This relatively temporary decline in long-term interest rates during the QE period may be one reason for the less persistent effects of quantitative easing.

One can also explain why quantitative easing has more immediate effects than conventional monetary policy, although it is difficult to provide evidence. On the one hand, the immediate response of output in the QE period can be explained by the convexity of the aggregate supply function, as also used above to interpret the stronger effects on output (and the weaker effects on prices). On the other hand, in order to explain the immediate response of prices, we can apply a cost channel theory. Here, the cost channel means that a decline in the short-term interest rate lowers the price level through cutting costs associated with, say, working capital, which has been suggested as one of the reasons why prices initially fall (i.e., creating a price puzzle) in response to an expansionary policy shock (see, e.g., Christiano et al. (1997) and Barth and Ramey (2001)). However, the influence of the cost channel would weaken under the circumstance where the short-term interest rate is near zero, which might cause earlier positive responses of prices to the expansionary policy shock in the QE period.

22 When estimating Models I and II, I could not obtain significant responses of prices and output for both periods. Also, note from Figure 6 that the macroeconomic effects of traditional monetary policy estimated with Model III become insignificant based on the 90 percent error bands. It is likely that this is because there are many parameters in the five-variable VAR for the relatively short sample and because an additional variable in each period has little helpful information to identify monetary policy shocks. For this reason, I treated the four-variable VAR as the empirical model in this paper rather than the five-variable VAR that has variables common to the two periods.

23 There are at least three reasons for the short-lived decline in long-term interest rates for the QE period. First, the BOJ's quantitative easing may not be strong enough to compress the long-term interest rates over longer periods. Second, in contrast to the first reason, the BOJ might succeed in raising inflation expectations. Third, the BOJ's massive purchases of Japanese government bonds might increase their default risk by loosening fiscal discipline.
3.3 Variance Decomposition Analysis

Finally, I compare the monetary policy effects based on the variance decomposition analysis. Figure 7 shows the results for the QE period obtained from the four-variable VAR including the spread. The variances of prices and output accounted for by monetary policy shocks become smaller when moving from Model I to Model III. More precisely, monetary policy shocks explain the variance of prices in the long run (at the horizon of the 60th month) by 47 percent in Model I, 24 percent in Model II, and only 7 percent in Model III, and the variance of output by 20 percent in Model I, 16 percent in Model II, and only 6 percent in Model III. These results are consistent with the results in the impulse response analysis, where an expansionary monetary policy shock has the largest effect in Model I and the smallest effect in Model III.

Figure 8 presents the results for the normal period. Comparison of Figures 7 and 8 shows qualitative and quantitative differences between the two periods. As a qualitative difference, the contributions of monetary policy shocks to both price and output fluctuations in the QE period reach a peak earlier than in the normal period, a similar feature to that in the impulse response analysis. As a quantitative difference, monetary policy shocks during the QE period make smaller contributions to the fluctuation in output, regardless of the model. This seems to contradict the results of the impulse response analysis, in which the effects of quantitative easing have larger effects on output than conventional monetary policy. This is partly because, in the QE period, money demand shocks account for the output fluctuation to some extent, whereas their contribution is very small in the normal period.

4 Conclusion

The Japanese economy has experienced a relatively longer period of unconventional monetary policy than many other major economies, which enables us to analyze its macroeconomic effects using a self-contained approach such as the VAR model. In this paper, I compared the macroeconomic effects of quantitative easing with those of conventional monetary policy in Japan, based on the sign-identified VAR model. In particular, I introduced three new approaches to this body of research: (1) including the long–short spread and defining an expansionary monetary shock as a shock that
not only increases bank reserves but also compresses the spread; (2) considering possible different
timings of the monetary policy effects by imposing magnitude and/or zero restrictions on the impact
responses of macroeconomic variables; and (3) combining the VAR model with the stock return
equation to estimate the size of a one-standard-deviation monetary policy shock in each period,
which makes possible a quantitative comparison of the two periods.

Several findings emerged from the analysis. First, I have shown that when comparing the effects
of monetary policy between the QE and normal periods, it is not necessarily appropriate to estimate
a VAR model that is common to both periods. For the QE period, the VAR model including the
long–short spread detects larger effects of quantitative easing than the VAR model with the short-
term policy rate that is also estimated for the normal period. In addition, significant effects of
quantitative easing are obtained in Models I and II (both allowing for the initial response of output
to a monetary policy shock), whereas significant effects of conventional monetary policy can be
found in Models II and III (the latter imposing zero restrictions on the initial responses of prices
and output).

Second, there are qualitative differences in monetary policy effects between quantitative easing
and conventional monetary policy. Overall, quantitative easing has immediate but less persistent
effects on both real economic activity and the price level. Moreover, the uncertainty concerning
the effects of quantitative easing on real economic activity is larger than that of the effects of
conventional monetary policy. This reflects conflicting results among previous studies on the actual
effects of quantitative easing in Japan.

Third, a quantitative comparison shows that quantitative easing has stronger effects on output
and weaker effects on the price level compared with conventional monetary policy. This result is
consistent with Gambacorta et al. (2014), explained by the convexity of the aggregate supply curve.

There are at least two limitations of this analysis, particularly regarding the estimation of the
effectiveness of quantitative easing. First, the analysis assumes that increases in bank reserves by
the BOJ have a constant effect over the QE period. However, it may be interesting to consider
possible changes in the effect over the three rounds of quantitative easing. Second, I only assessed
the implementation effects of quantitative easing. However, the announcement effects regarding its
introduction, extension, taper, and exit may be comparable to or even larger than the implementation effects. Identifying the announcement shocks along with the implementation shocks should be a topic for future research.

References


Appendix A. Data Description

This appendix provides detailed descriptions of the data.

**Prices** \((p)\)

The price level is measured by the consumer price index (all items less fresh food), which is on the website of the Statistics Bureau, Ministry of Internal Affairs and Communications in Japan (http://www.stat.go.jp/english/). Because the seasonally adjusted series is only available from January 2005, I apply the seasonal adjustment method of the US Census Bureau’s X-12 to the original non-seasonally adjusted series for the period January 1970 to March 2014.

**Output** \((x)\)

Output is measured by the seasonally adjusted industrial production index, which is provided by the Ministry of Economy, Trade and Industry in Japan on its website (http://www.meti.go.jp/english/).

**Short-term interest rates** \((r_{\text{short}})\)

Following Miyao (2002), the series for the call rate (the Japanese short-term interbank rate and the BOJ’s policy rate) is constructed by connecting the uncollateralized call rate (overnight, monthly average of daily observations) from July 1985 to the collateralized call rate (overnight, monthly average of daily observations) before that time. As the uncollateralized rate is higher than the collateralized rate by the amount of the risk premium, Miyao (2002) added a mean difference between the two rates to the collateralized rate series. I adopt this approach and calculate the mean
difference over the period from July 1985 to June 1995, resulting in a risk premium of 13 basis points. The two series for the call rate are available from the BOJ’s website (http://www.boj.or.jp/en/) with the series code of ST’STRACLUCON for the uncollateralized rate and ST’STRACLCVOON for the collateralized rate.

**Spread ($\Delta$)**

The long–short spread is constructed by subtracting the call rate from the 10-year yield on Japanese government bonds. The latter daily data (series code; LBSHIYNK) is from Nikkei Media Marketing Inc., and is then transformed into a monthly average series.

**Bank reserves ($m$)**

Bank reserves are the monthly average outstanding balances adjusted for the reserve requirement rate changes, which are taken from the BOJ’s website with the series code of BJ’MABS1AN125. I use the X-12 seasonal adjustment method over the period from January 1970 to March 2014.

**Stock returns ($\Delta s$)**

The stock return in a month is the log-difference (multiplied by 100) between the stock price index at the end of the current month and that at the end of the previous month. As a measure of the stock price index, I use the *Nikkei Stock Average* published by Nikkei Inc. on its website (http://indexes.nikkei.co.jp/en/nkave/).

**Appendix B. Details of the Identification in Models I to III**

This appendix describes how to impose the impact restrictions of Models I to III proposed in Section 2.4. Let $A$ denote the Cholesky factor of the variance–covariance matrix $\Sigma$ of reduced-form VAR errors, and let $U$ denote the orthogonal matrix obtained by drawing a $4 \times 4$ matrix $K$ from the $N(0,1)$ distribution, taking the $QR$ decomposition of $K$ such that $K = Q \cdot R$, and defining $U = Q’$. Then, the structural impact matrix for Model I is computed as $\tilde{A}_I = A \cdot U$. To accomplish the identification of Model I, we need to check whether the resulting impulse response functions (their
impact responses corresponding to the impact matrix \( \tilde{A}_I \) satisfy the two magnitude restrictions described in Section 2.4, as well as all the sign restrictions shown in Table 1.

Next, to impose the zero restriction of Model II, I apply the method used in Baumeister and Benati (2013).24 More precisely, I define a rotation matrix \( H_{II} \) as

\[
H_{II} = \begin{bmatrix}
\cos(\varphi) & 0 & 0 & \sin(\varphi) \\
0 & 1 & 0 & 0 \\
0 & 0 & 1 & 0 \\
-\sin(\varphi) & 0 & 0 & \cos(\varphi)
\end{bmatrix}
\]

with \( H_{II} \cdot H_{II}' = I_4 \), where \( I_4 \) is a 4 \times 4 identity matrix. The rotation angle \( \varphi \) is defined as \( \varphi = \tan^{-1}(\tilde{A}_{1,4}/\tilde{A}_{1,1}) \), where \( \tilde{A}^{i,j} \) denotes the \((i,j)\) element of \( \tilde{A} \) and \( \tilde{A} \) is defined as \( \tilde{A} = A \cdot U \).

Then, I obtain the impact matrix for Model II as \( \tilde{A}_{II} = \tilde{A} \cdot H_{II} \) that has a zero in the (1, 4) position, which means a zero initial response of prices to a monetary policy shock.

Finally, to achieve the identification of Model III, I rely on the approach taken by Gambacorta et al. (2014). Define a rotation matrix \( H_{III} \) as

\[
H_{III} = \begin{bmatrix}
0 & 0 & 0 & 0 \\
0 & 1 & 0 & 0 \\
0 & 0 & \cos(\theta) & -\sin(\theta) \\
0 & 0 & \sin(\theta) & \cos(\theta)
\end{bmatrix}
\]

with \( H_{III} \cdot H_{III}' = I_4 \). The rotation angle \( \theta \) is uniformly drawn from \([0, \pi]\). Then, the impact matrix for Model III is defined as \( \tilde{A}_{III} = A \cdot H_{III} \), whose (1, 4) and (2, 4) positions are zero, implying that impact responses of both prices and output to a monetary policy shock are constrained to be zero.25

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24Baumeister and Benati (2013) impose a zero restriction on the impact response of the policy rate to the spread shock.

25\( \tilde{A}_{III} \) has an additional three zeros in the (2, 1), (3, 1), and (3, 2) positions. In particular, the zero in the (2, 1) position means that an output shock has no effect on prices at impact, which violates one of the sign restrictions in Table 1. However, this difference is minor in that no immediate response of prices to the output shock is not an implausible restriction.
Table 1. Identifying Sign Restrictions

<table>
<thead>
<tr>
<th>Shock Variable</th>
<th>VAR model (1)</th>
<th>r</th>
<th>m</th>
<th>Δs</th>
</tr>
</thead>
<tbody>
<tr>
<td>( x )</td>
<td>( \epsilon_t )</td>
<td>( (h=0) )</td>
<td>+</td>
<td>( (h=0) )</td>
</tr>
<tr>
<td>( p )</td>
<td>( (h=0,\ldots,5) )</td>
<td>+</td>
<td>( (h=0) )</td>
<td>+</td>
</tr>
<tr>
<td>( \Delta s )</td>
<td>( \epsilon_{x,t} )</td>
<td>*</td>
<td>( (h=0) )</td>
<td>+</td>
</tr>
<tr>
<td>( \epsilon_{p,t} )</td>
<td>( (h=0) )</td>
<td>-</td>
<td>( (h=0) )</td>
<td></td>
</tr>
<tr>
<td>( \epsilon_{mp} )</td>
<td>( (h=0) )</td>
<td>-</td>
<td>( (h=0) )</td>
<td></td>
</tr>
</tbody>
</table>

Notes: \( r \) denotes either the short-term interest rate or the spread. An \( '*' \) means that an additional restriction is imposed on the impact response (see Section 2.4). A blank entry indicates that no restrictions are imposed.
Table 2. Stock-Returns Responses to Expansionary Monetary Policy Shock

<table>
<thead>
<tr>
<th></th>
<th>(1) Median of $\beta_4$ for the QE period</th>
<th>(2) Median of $\beta_4$ for the normal period</th>
<th>Ratio (2)/(1)</th>
</tr>
</thead>
<tbody>
<tr>
<td>Model I</td>
<td>0.1986</td>
<td>0.1805</td>
<td>0.9089</td>
</tr>
<tr>
<td>Model II</td>
<td>0.1768</td>
<td>0.2539</td>
<td>1.4361</td>
</tr>
<tr>
<td>Model III</td>
<td>0.1311</td>
<td>0.1988</td>
<td>1.5164</td>
</tr>
</tbody>
</table>
Figure 1. The Call Rate and Bank Reserves

Notes: The solid vertical lines represent the end of the normal period (June 1995) and the beginning of the QE period (March 2001). The dashed vertical lines represent the end of the first QE (March 2006), the beginning of the second QE (October 2010), and the beginning of the third QE (April 2013).
Figure 2. Impulse Response Functions to an Expansionary Monetary Policy Shock
(The VAR Model with the Call Rate; the QE Period)

Notes: The solid, dashed, and dotted lines indicate the median responses, the 16th and 84th percentile error bands, and the 5th and 95th percentile error bands, respectively.
Figure 3. Impulse Response Functions to an Expansionary Monetary Policy Shock
(The VAR Model with the Spread; the QE Period)

Notes: See notes in Figure 2.
Figure 4. Impulse Response Functions to an Expansionary Monetary Policy Shock
(The VAR Model with the Call Rate: the Normal Period)

Notes: See notes in Figure 2.
Figure 5. Comparison of the Macroeconomic Effects of Monetary Policy

Notes: The solid and dashed lines indicate the median and the 16th and 84th percentile error bands for the QE period (corresponding to Figure 3). The shaded areas represent the regions within the 16th and 84th percentile error bands for the normal period (corresponding to Figure 4).
Figure 6. Impulse Response Functions to an Expansionary Monetary Policy Shock
(The Five-Variable VAR Model with the Call Rate and the Spread: Model III)

Notes: See notes in Figure 2. The right column shows impulse responses of the implied long-term interest rate.
Figure 7. Variance Decomposition
(The VAR Model with the Spread; the QE Period)
Figure 8. Variance Decomposition
(The VAR Model with the Call Rate; the Normal Period)